

## Macroeconomics and agriculture in Tunisia

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## MACROECONOMICS AND AGRICULTURE IN TUNISIA

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MACROECONOMICS AND AGRICULTURE IN TUNISIA

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**Abstract**

This paper aims to analyse the impact of changes in the monetary policy and the exchange rate on agricultural supply, prices and exports. The methodology used is based on the multivariate cointegration approach. Ten variables are considered: interest and exchange rates, money supply, inflation, agricultural output and input prices, agricultural supply and exports, income and the rate of commercial openness. The sample period covers annual data from 1967 to 2002. Due to the short sample period, two subsystems are considered. First, long-run relationships are identified in each subsystem. Second, both subsystems are merged in order to calculate the short-run dynamics. The results indicate that changes in macroeconomic variables have an effect on the agricultural sector but the reverse effect does not hold.

**Key words:** Macroeconomic policy, agro-food sector, Tunisia, Africa, dynamic relationships

# MACROECONOMICS AND AGRICULTURE IN TUNISIA

## Abstract

This paper aims to analyse the impact of changes in the monetary policy and the exchange rate on agricultural supply, prices and exports. The methodology used is based on the multivariate cointegration approach. Ten variables are considered: interest and exchange rates, money supply, inflation, agricultural output and input prices, agricultural supply and exports, income and the rate of commercial openness. The sample period covers annual data from 1967 to 2002. Due to the short sample period, two subsystems are considered. First, long-run relationships are identified in each subsystem. Second, both subsystems are merged in order to calculate the short-run dynamics. The results indicate that changes in macroeconomic variables have an effect on the agricultural sector but the reverse effect does not hold.

## 1. Introduction

The ongoing globalisation process in the world economy is a big challenge for Tunisia, a country which has suffered a complex process of structural economic reforms. The Adjustment Structural Program, implemented in 1986, generated a new environment of economic success. All sectors of the economy started to recover and exports dramatically increased, being one of the main contributors to economic development. In the last five years the Tunisian GDP increased at a 5.5% annual rate while inflation was maintained around 3.5 %.

The agro-food sector in Tunisia plays an important role in the Tunisian economy. It generates around 14% of the total GDP, employs 22% of the total labour force and agro-food exports represent around 15% of total exports, although they still depend to a great extent on weather conditions. Moreover, since 1986, the agricultural sector has been undergoing a modernization process characterized by a progressive intensification and the use of technology. However, the agricultural production has not been able to meet the needs of an increasing population. In general, the Government favoured imports of raw materials and food, which have provoked a progressive deterioration of the trade balance. The agricultural policy was, then, oriented in two directions: 1) to promote the production of agricultural products in which self-sufficiency was low, through the implementation of a subsidies program (food security); and 2) to encourage the production of food products in which Tunisia had traditionally had a competitive advantage (olive oil, fruit, vegetables, etc) to finance the agricultural trade deficit.

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In many cases, the results from such policies were, to some extent, different from those expected as the effect of many macroeconomic variables (as a consequence of the Adjustment Structural Program) were not taken into account. Although not explicitly recognized, changes in the macroeconomic policy have become increasingly important for the agro-food sector, as Tunisian agriculture has become more capitalized, more dependent on international markets and, thus, more vulnerable to changes in interest rates, exchange rates and international growth rates.

The aim of this paper is, precisely, to provide a methodological approach, taking data limitations into account to explain the relationships between macroeconomic variables and the agricultural sector in Tunisia. Special attention is paid to the distinction between long-run structural relationships and short-run dynamics. As far as we know, this is the first attempt to analyse such relationships in Tunisia. The existing literature on Tunisia is quite descriptive, focussing on the evolution of agricultural trade flows which are only explained by changes in the agricultural policy and exchange rates (Arfa, 1994; Ben Said, 1994 Allaya, 1995; and El Abassi, 1995, among others).

Since the mid seventies, a number of theoretical and empirical studies have analysed the impact of macroeconomic variables on the relative performance of the agricultural sector (see In and Mount, 1994, for a review of the literature on this topic). In the early studies, macroeconomic variables (income, interest rate, exports,...) were introduced as purely exogenous in agricultural sector models. The paper by Schuh (1974) could be considered as the starting point of a second group of studies emphasizing the role of the exchange rate in explaining agricultural variable fluctuations (Chambers and Just, 1979, 1981; Longmire and Morey, 1983; and Batten and Belongia, 1986). However, these empirical investigations neglect not only the possible effect of exchange rate changes on other macroeconomic variables (which can influence agricultural prices and exports indirectly) but also the effects of other macroeconomic variables (such as interest rates) both on the exchange rate and on agricultural variables. In this context, Chambers (1984) develops a general equilibrium model in order to analyse the effect of macroeconomic variables on agricultural trade where not only the exchange rate, but income, the interest rate, as well as the usual agricultural variables, are considered as endogenous.

Finally, it is possible to identify a third group of papers dealing with the analysis of the dynamic linkages between monetary variables and the agricultural sector. The

main issue is whether levels of agricultural and non-agricultural prices respond proportionally to changes in the level of money supply in the long run, and whether money is neutral in the short run. The question of money neutrality in the agricultural sector, as well as the speed of price adjustments, have been considered of central importance for policy analysis (Bordo, 1980; Tweeten, 1980, Bessler and Babula, 1987; Devadoss and Meyers, 1987; Taylor and Spriggs, 1989; Robertson and Orden, 1990; Larue and Babula, 1994; Dorfman and Lastrapes, 1996; Loizou et al., 1997; Kargbo, 2000, among others). Bordo (1980) argues that agricultural commodities tend to be more highly standardised and therefore exhibit lower transaction costs than manufactured goods. Consequently, agriculture is characterised to have rather short term contracts, which lead a faster response to a monetary shock. Alternatively, Tweeten (1980) argues that price shocks stemming in oligopolistic non-agricultural sector and accommodated by expansionary monetary policy, cause inflation and place agricultural in a price-cost squeeze.

Results from most of the above-mentioned studies differ substantially from each other and, in many cases, are even contradictory. There are alternative explanations for such differences: the samples are not homogeneous, the number of variables included differs as well as their treatment as endogenous or exogenous, and they use different methodological approaches. Moreover, as Ardeni and Freebairn (2002) pointed out, many studies lack an appropriate treatment of the time series properties of data implying misleading results especially in the case of earlier research. However, there seems to be a consensus on the fact that models analysing macroeconomic linkages to the agricultural sector should include the more relevant macroeconomic variables of the country being analysed and should treat them as endogenous (Devadoss et al., 1987; Taylor and Spriggs, 1989; Denbaly and Torgerson, 1991; Thraen et al., 1992; Devadoss and Chaudhary, 1994; In and Mount, 1994; Ben Kaabia and Gil, 2000; Ivanova et al., 2003, among others). Partly for this reason, most of the analyses on this topic have recently been conducted using Vector Autoregression (VAR) models. This is also the methodological approach we have followed in this paper although adapted to take data limitations and their stochastic properties into account.

The paper is organized as follows. The data used in this study, their stochastic characteristics and the methodological approach are presented in Section 2. Long-run

equilibrium relationships are analysed in Section 3. The short-run dynamics are considered in Section 4. Finally, some concluding remarks are outlined.

**2. Data and methodological approach**

Since Sims’ (1986) seminal paper, VAR models have been one of the most widely used tools to analyse the dynamic relationships between macroeconomic and agricultural variables. In VAR models, all variables are considered endogenous and no zero/one restrictions are imposed on the variables in the system. Moreover, it is possible to calculate the short-run responses to a shock in one variable in the system from any other variable, by offering a convenient way to characterise data without involving economic theory to restrict the dynamic relationships among variables. Cooley and LeRoy (1985), among others, have criticised the usefulness of such an atheoretical approach for policy analysis. To overcome this problem, "Structural" VAR (SVAR) models have been used (Bernanke, 1986, Sims, 1986, and Blanchard and Quah, 1989) which allow the researcher to specify and test restrictions based on economic theory prior to calculating the impulse response functions (Orden and Fackler, 1989). However, the economic theory driving the restrictions is “weak”; although the identifying restrictions imposed are consistent with economic theory, they have not been derived from fully specified economic models (see Cooley and Dwyer, 1998). To resolve this dilemma, Pesaran and Shin (1998) have proposed the use of generalised impulse response functions to compute short-run dynamics for a set of variables, which, unlike the traditional impulse response analysis, is invariant with respect to the ordering of those variables.

Finally, recent developments in time series analysis have modified the econometric framework for analysing the relationships between macroeconomic variables and the agricultural sector. The concepts of non-stationarity and cointegration have become very popular and have to be explicitly tested to properly specify an econometric model. In this new context, Johansen (1988) and Johansen and Juselius (1990, 1992 and 1994) provide an interesting methodology that allows the researcher to distinguish between the short and the long run. On the one hand, it is possible to identify the long-run structural relationships among a set of variables and how variables in the system adjust to deviations from such long-run equilibrium relationships. On the



other hand, it is possible to calculate the impulse response functions in a similar way to that in the VAR models. This distinction is useful as economic restrictions are considered to be long-run in nature while it is also interesting, for policy analysis, to know how the system adjusts to disequilibrium.

In this paper we have followed this methodological approach although we have introduced some modifications in order to adapt it to the data limitations. Availability of data is a major problem for economic modelling in Tunisia. It is difficult to find a large enough sample period for many economic variables. In this study, 10 variables have been considered which collect the most important information related to macroeconomic variables and the agricultural sector (see the Appendix for data sources and units of measurement):

- a) Real exchange rate (ER), defined as national currency (TND) per US dollar taking into account both the US and Tunisian consumer price indices<sup>1</sup>;
- b) Real money supply<sup>2</sup> (M) (money supply ( $M_2$ ) divided by the consumer price index);
- c) Interest rate (R), defined as the one-year money market interest rate;
- d) Inflation (P) expressed as the Consumer Price Index in first differences;
- e) Real Gross Domestic Product (GDP);
- f) Real farm output prices (PP), calculated as nominal farm output prices divided by the Consumer Price Index);
- g) Real farm input prices (IP)<sup>3</sup>, calculated as nominal farm input prices divided by the Consumer Price Index);

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<sup>1</sup> A multilateral rather than a bilateral real exchange rate would have provided better information about Tunisia's competitiveness in the trade market and on agricultural exports dynamics. However, these data have only been available since 1983. In any case, when comparing the evolution of both rates since the information has been available, there do not seem to be any significant differences in relation to trends and turning points.

<sup>2</sup> The objective of the monetary policy in Tunisia has been to keep inflation close to that of its main competitors. Traditionally, the government establishes the growth rate of the money supply ( $M_2$ ) 2% lower than the expected growth of the GDP. However, this objective has always been subject to revision within the year.

<sup>3</sup> In this study the "Index de pris de vente d'engrais" has been used as an aggregate input price index was not available for the whole sample. This index comprises most of the intermediate inputs used in agriculture and is published by the Institut National de la Statistique of Tunisia. The same index was used



- h) Real agricultural exports<sup>4</sup> (AX), calculated as the nominal exports value divided by the Consumer Price Index);
- i) Agricultural output (AP), calculated as the value of the Tunisian Agricultural Output divided by the Consumer Price Index; and
- j) Rate of commercial openness (RCO) calculated by dividing the international trade flows (imports + exports) by the GDP. This variable provides an indication on how the Tunisian economy is inserted into the world trade.

All the variables are in logarithms, except for the interest rate and inflation, which are in a percentage form and are divided by one hundred to make the estimated coefficients comparable with the logarithmic changes. The sample period covers annual data from 1967 to 2002. Time series univariate properties have been examined by using unit root tests. As in small samples such tests have limited power (Blough, 1992), two alternative unit root tests developed by Elliot et al., (1996) and Ng and Perron (2001) as well as the stationary test from Kwiatkowski et al. (1992) (KPSS) have been applied. All the tests indicated that all the variables were  $I(1)$ <sup>5</sup>.

Moreover, in order to check the impact of the implementation of the 1986 Structural Adjustment Program on the evolution of the above mentioned variables, unit root test allowing for the presence of structural breaks have been performed. From the seminal work by Perron (1989) on the influence of structural breaks on unit root tests, much of the literature have dealt with the case in which a break occurs during one period only. However, it may be more reasonable to think that breaks occur over a number of periods and display smooth transition to a new level. Saikkonen and Lütkepohl (2002) and Lanne et al. (2002) develop such a model which adds to the deterministic term shift functions of a general nonlinear form  $f_t(\Phi)$ . In a model with a linear trend term and shift the underlying model is given by:

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by Dhehibi and Lachaal (2006) to calculate the sources of agriculture productivity growth in Tunisia during the period 1961-2000.

<sup>4</sup> Agricultural exports in Tunisia mainly refer to olive oil (40% of total agricultural exports), fish (20%); dates (10%) and citrus fruit (3%). Around 60% is exported to the European Union (EU). Except for the dates, the EU set a maximum amount to be imported with a lower tariff. Only in the case of the olive oil, in some years total exports have exceeded the maximum amount allowed. Theoretically, the Tunisian government does not directly subsidise exports. However, from the 60's to the mid 80's agricultural policy was based on guaranteed prices and subsidies for agricultural inputs for selected agricultural products (mainly food staples but also for some exporting products).

<sup>5</sup> Results are not shown due to space limitations. They are available upon request.

$$y_t = \mu_0 + \mu_1 t + f_t(\Phi)' \gamma + z_t \quad (1)$$

where  $\Phi$  and  $\gamma$  are unknown parameters and  $z_t$  are residual errors. The shift function  $f_t(\Phi)$  is able to characterize whether the break is abrupt and complete within one time period or more gradual. For instance, consider the simple shift function with break at time  $T_B$ :

$$f_t = d_{it} = \begin{cases} 0, & t < T_B \\ 1, & t \geq T_B \end{cases} \quad (2)$$

In this case, the structural break is represented by a simple dummy variable with shift date  $T_B$ . On the other hand, a more general shift function which allows for sharp, one-time shifts and a more gradual shift to a new level beginning at time  $T_B$  can be expressed as:

$$f_t(\Phi) = \begin{cases} 0, & t < T_B \\ 1 - \exp\{-\Phi(t - T_B + 1)\}, & t \geq T_B \end{cases} \quad (3)$$

As can be observed, in this latter case, when  $\Phi \rightarrow \infty$ , the shift function becomes a dummy variable. Saikkonen and Lütkepohl (2002) and Lanne et al. (2002) propose unit root tests for the model (1) in first differences which are based on estimating the deterministic term first by a generalized least squares (GLS) procedure under the unit root null hypothesis and subtracting it from the original series. Then an Augmented-Dickey-Fuller (ADF) type test is performed on the adjusted series, which also includes terms to correct for estimation errors in the parameters of the deterministic part. As in the case of the ADF statistic, the asymptotic null distribution is nonstandard. Critical values are tabulated in Lanne et al. (2002).

The user of the test has to decide on the AR order and the shift date. If the latter is known, the desired shift function may be included and the AR order may be chosen in the usual way with the help of order selection criteria, sequential tests and model checking tools. This approach is extended to a situation of unknown break date by Lanne et al. (2003), which is the approach followed in this paper. Results are shown in Table 1 and indicate that the null of unit root against the alternative of stationarity with a structural break can not be rejected. Finally, it is important to note that the endogenously determined structural breaks for each variable are different and it seems that the implementation of the Structural Adjusted Program has not generated an abrupt change in the evolution of the analyzed series.

Taking into account the number of variables, the number of observations available for each variable and that all variables are  $I(1)$ , the methodological approach followed in this paper consisted of the following steps:

- i) The ten-variable system is divided into two subsystems. The first one has been defined by including: the real money supply, inflation, the GDP, the farm input and output prices and the interest rate. Furthermore, taking into account the characteristics of the Tunisian economy, we have considered the interest rate as purely exogenous. The second subsystem includes the following seven variables: the farm input and output prices, the agricultural exports, the agricultural production, the exchange rate, the interest rate and the rate of commercial openness. Within this subsystem, the interest rate and the rate of commercial openness are defined as purely exogenous<sup>6</sup>.
- ii) Under the assumption of exogeneity for certain variables, the multivariate cointegration procedure developed by Pesaran et al. (2000) is used to test for cointegration in both subsystems. Moreover, cointegration vectors are identified as long-run meaningful economic relationships.
- iii) Merging the results from the two subsystems into a single system with the original 10 variables, impulse response functions are computed to analyse the short-run dynamics and to test the exogeneity assumptions made in the first step.

(Insert Table 1)

**3. Long-run analysis**

*3.1. Model specification and cointegration rank*

All the variables in each subsystem were  $I(1)$  and, so, a Vector Error Correction Model has been specified for each subsystem. The methodology developed by Pesaran et al. (2000) is used to determine the cointegration rank. These authors modified the Johansen (1988) procedure to explicitly allow for the introduction of exogenous variables. The base-line econometric specification for multivariate cointegration is a

<sup>6</sup> In a further sep in the modelling process, specific tests will be carried out to test for the exogeneity of the mentioned variables.

VAR(p) representation of a  $k$ -dimensional time series vector  $Y_t$  reparametrized as a Vector Error Correction Model (VECM):

$$\Delta Y_t = \mu D_t + \Gamma_1 \Delta Y_{t-1} + \dots + \Gamma_{p-1} \Delta Y_{t-p+1} - \Pi Y_{t-1} + e_t \quad (4)$$

where,  $Y_t$  is a  $(k \times 1)$  column vector of variables;  $D_t$  is a vector of deterministic variables (intercepts, trend...); and  $\mu$  is the matrix of parameters associated with  $D_t$ ;  $\Gamma_i$  are  $(k \times k)$  matrices of short-run parameters ( $i=1, \dots, p-1$ ), where  $p$  is the number of lags;  $\Pi$  is a  $(k \times k)$  matrix of long-run parameters and  $e_t$  is the vector of disturbances  $niid(0, \Sigma)$ .

When exogenous variables are considered, the  $Y_t$  vector can be partitioned as  $Y_t = (Z_t', X_t')'$ , where  $Z_t$  is an  $(m \times 1)$  vector of endogenous variables and  $X_t$  is an  $(n \times 1)$  vector of exogenous variables ( $n=k-m$ ), which can be considered as the “long-run forcing” variables in the system, that is, changes in  $X_t$  have a direct influence on the variables  $Z_t$ , while they are not affected either by the changes in the equilibrium relationships nor by past changes in  $Z_t$ . This is equivalent to the notion that the set of variables  $Z_t$  do not Granger-cause  $X_t$ .

Under such circumstances, Pesaran et al. (2000) show that the  $k$ -variable system defined in (4) can be decomposed to following two subsystems:<sup>7</sup>.

- Conditional subsystem:  $\Delta Z_t = \delta D_t + \Lambda \Delta X_t + \sum_{i=1}^{p-1} \Psi_i \Delta Y_{t-i} + \Pi_z Y_{t-1} + u_t \quad (5)$

- Marginal subsystem:  $\Delta X_t = \mu_x + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta X_{t-i} + \varepsilon_{xt} \quad (6)$

“Variables cannot be exogenous per se” (Hendry, 1995). A variable can only be exogenous with respect to a set of parameters of interest. Hence, if the variables  $X_t$  are deemed to be exogenous with respect to parameters in (5), the marginal model (6) can be neglected and the conditional model (5) is complete and sufficient to sustain valid inference. Hence, the knowledge of the marginal model will not significantly improve the statistical or forecasting performance of the conditional model.

Following this line of reason, the conditional model in equation (5) is used to test for cointegration, which is equivalent to testing for the Rank ( $r$ ) of matrix  $\Pi_z$ . So, the

<sup>7</sup> Under this decomposition, variables in  $X_t$  are assumed to be weakly exogenous with respect to the cointegration space. Moreover, if the variables in  $Z_t$  do not Granger-cause  $X_t$ , then these variables are assumed to be strongly exogenous with respect to the cointegration space, that is, they would be only explained by their own past in the marginal subsystem.

null hypothesis to test that there exists at least  $r$  cointegrating vectors among the variables included in each subsystem can be formulated as:

$$H_r : \text{Rank} [\Pi_z] = r \qquad r=0, \dots, m \qquad (7)$$

To test for the number of cointegrating vectors ( $r$ ), Pesaran et al., (2000), following Johansen (1988), proposed two statistics: the trace statistic (which tests whether there are at least  $r$  cointegrating vectors against the maintained hypothesis; and the  $\lambda_{\max}$  statistic (which tests whether there are  $r$  cointegrating vectors against the alternative  $r=r+1$ ). If the hypothesis of cointegration is not rejected ( $0 < r < m$ ),  $Y_t$  is said to be cointegrated in the sense that there exists a  $k \times r$  matrix  $\beta$  such that  $(\beta'Y_{t-1})$  is stationary and, consequently, the cointegration relationships can be formally expressed as  $\Pi_z = \alpha_z \beta'$ . Each column of matrix  $\beta$  represents a cointegrating vector, whereas the rows of matrix  $\alpha_z$  represent the adjustment coefficients that determine the speed of adjustment of the  $m-1$  equations to disequilibrium.

The procedure outlined above has been applied to the two subsystems described in the last section. However, in empirical applications, the choice of  $r$  is frequently sensitive to: i) the deterministic terms included in the system (such as a constant and/or a trend) and on the way in which such components interact with the error correction term; and ii) the appropriate lag length to ensure that the residuals are Gaussian. In this paper, both subsystems are estimated including two lags<sup>8</sup> and a constant restricted to the cointegration space<sup>9</sup>. Multivariate tests for autocorrelation (Godfrey, 1988) and normality (Doornik and Hansen, 1994) have been carried out to check for model statistical adequacy before applying the reduced rank tests. Results indicated that both subsystems could be considered correctly specified<sup>10</sup>.

Table 2 shows the results of the cointegration tests in both subsystems. As can be observed, for the first subsystem (upper part of Table 2) the results of the  $\lambda$ -max and the

<sup>8</sup> A small-sample adjusted Likelihood Ratio statistic has been used considering a maximum lag of three periods.

<sup>9</sup> Results from unit root tests indicated that almost all the variables were non-stationary with no-zero means.

<sup>10</sup> Results from multivariate first-order autocorrelation tests were 21.14 and 23.85 for the first and the second subsystem, respectively, which were well below the critical value at the 5% level of significance ( $\chi^2_{25} = 37.65$ ). Results from multivariate normality tests were 14.43 and 17.83, for the first and the second subsystem, respectively, which were well below the critical value at the 5% level of significance ( $\chi^2_{10} = 18.31$ ).

trace tests indicate that there are two cointegration vectors among the six variables included while for the second subsystem (lower part of Table 2) the results differ depending on the level of significance (two and three cointegration vectors for the 5 and 10% levels of significance, respectively).

(Insert Table 2)

Taking into account the relatively large dimension of the VECM and the small sample available, the outcome of the test procedure has to be interpreted with some caution. Several simulation studies show (Abadir et al., 1999; Gredenhoff and Jacobson, 2001; and Johansen and Juselius, 2000) that the asymptotic critical values may not be very close approximations in small samples. Because of this, we have also studied the roots of the companion matrix and the t-ratios of the  $\alpha_z$  parameters from the last cointegration vector (Juselius, 1995). For both subsystems, all the roots were inside the unit circle, indicating that all the variables were I(1). Moreover, the eigenvalues of the companion matrix show that, for both subsystems, the first four roots were close to unity while the rest were quite small. In other words, we could not reject the null of two and three cointegrating vectors for the first and the second subsystems, respectively. Finally, all the t-statistics of the  $\alpha_z$  parameters of the third cointegration vector for the first subsystem were not significant, while in the second subsystem, some of them were significant<sup>11</sup>. Thus, the first subsystem has been specified with two cointegrating vectors, whereas three cointegration vectors have been chosen for the second one.

### 3.2 Long-run structural relationships

The estimated  $\beta$  and  $\alpha_z$  parameters are presented in Table 3, where  $\beta$  is presented in normalized form. As can be observed, all the parameters of the long-run equilibrium relationships found in each subsystem are statistically significant at the 5% level of significance. Identifying economically interpretable relations is the primary aim of this analysis. However, Juselius (1994) argues "the interpretation of the unrestricted cointegration space is far from straightforward when there are more than one cointegrating vector". Moreover, Johansen and Juselius (1994) suggest that only sometimes can the unrestricted cointegrating vectors, surprisingly, be directly interpreted in terms of theoretical economic relationships. Thus, some restrictions are needed in order to obtain a structural representation of such relationships.

<sup>11</sup> Results are not presented due to space limitations but are available from the authors upon request.

(Insert Table 3)

First subsystem

Taking into account the variables included in the model as well as the economic theory which relates those variables, the following hypothetical cointegration relations could be expected:

i) A money demand equation in real terms in which the monetary aggregate is related to inflation in Tunisia, the Gross Domestic Product and an opportunity cost represented by the interest rate:

$$(\beta_1^{sys1})' Y_t : RM_t = \beta_{GDP}^1 GDP_t + \beta_R^1 R_t + \beta_P^1 P_t + \mu^1 + \varepsilon_{1t} \quad (8)$$

It is expected that  $\beta_{GDP} > 0$ ;  $\beta_R < 0$  and  $\beta_P < 0$ . If  $\beta_{GDP} = 1$ , equation (8) would be consistent with the Quantity Theory of Money, whereas  $\beta_P = 0$  would exclude inflation playing a role in the demand for money in Tunisia.

ii) A price transmission equation:

$$(\beta_2^{sys1})' Y_t : IP_t = \beta_{PP}^2 PP_t + \mu^2 + \varepsilon_{2t} \quad (9)$$

from which it is possible to test the homogeneity condition:  $-\frac{\beta_{IP}}{\beta_{PP}} = \frac{\partial IP}{\partial PRC} \frac{PP}{IP} = 1$

The two equations can be written more compactly as:  $\beta_{sys1}' Y_{t-1} = \varepsilon_t \sim I(0)$

$$\text{where: } \beta'^{sys1} = \begin{bmatrix} 1 & * & * & 0 & 0 & * & * \\ 0 & 0 & 0 & 1 & -1 & 0 & * \end{bmatrix} \quad (10)$$

In this paper, a two-step procedure is going to be used in order to check if (10) is supported by the data. In the first step, each single restricted relation (8)-(9) is tested for stationarity, leaving the other relations unrestricted. In other words, to test whether the restrictions imposed are compatible with a stationary relationship. The second step involves jointly considering the full identification of the two relationships. Juselius (1998) points out that this approach maximizes the chance of finding a correct full identification of long-run relations.



Hypotheses related to the first step adopt the general form  $H_{0i}: \beta=(H_i\phi,\omega)^{12}$ . In such an expression, the restrictions to be tested are only placed in a single cointegration vector while the remaining  $(r-1)$  vectors are considered unrestricted. Johansen and Juselius (1992) suggest that this test can be used when we wish to test if there is any vector in the cointegration space that linearly combines the variables in a particular hypothesized stationary relationship. Several hypotheses have been considered and tested. The specification of such hypotheses, as well as the main results found, is shown in Table 4. With respect to the first relationship, three different hypotheses have been tested. In the first one ( $H_{01}^{sys1}$ ), it is tested whether real money is cointegrated with the interest rate, GDP and inflation, imposing income homogeneity as well. Results from the Likelihood Ratio (LR) statistic indicate that the null cannot be rejected<sup>13</sup>. In the second hypothesis ( $H_{02}^{sys1}$ ), an additional restriction is considered ( $\beta_p^1=0$ ). This hypothesis is strongly rejected, which means that the monetary authority is not fixing the monetary policy by taking an aggregate money stock into account. The third hypothesis ( $H_{03}^{sys1}$ ) is similar to the first one but excluding income homogeneity. Also in this case, we fail to reject the null hypothesis. Finally, with respect to the second relationship, the hypothesis ( $H_{04}^{sys1}$ ) tests for price homogeneity in the agricultural sector. The LR statistic is under the critical value suggesting that monetary policy has a neutral effect on the real food-based prices. This means that, in the long run, input prices and output prices react in the same way and magnitude to changes in money supply.

Once it has been checked that each single equation is a cointegrated relationship, the second step consists of testing a full identification of the structural long-run relationships following Johansen and Juselius (1994). Taking the results shown in the upper part of the Table 4 into account, two hypotheses have been tested. The first one ( $H_{05}^{sys1}$ ) jointly tests hypotheses  $H_{01}^{sys1}$  and  $H_{04}^{sys1}$ , whereas the second tests hypotheses

<sup>12</sup> See Johansen and Juselius (1992) for a full description of the procedure to formulate and test such hypotheses.

<sup>13</sup> Several authors such as Reimers (1992) and Abadir et al. (1999) pointed out the tendency of likelihood ratio tests to over-reject in small samples when testing for the cointegration rank. In addition, some simulation studies have shown that, in small samples, the use of the  $\chi^2$  critical values can generate considerable size distortions when testing for hypotheses on the cointegration space (Gredenhoff and Jacobsen, 1998). Garratt et al. (1999) used a bootstrapping exercise to obtain critical values for testing the over-identification restrictions. The resulting critical values were higher than the asymptotic ones. This result would imply that the over-identification restrictions tested here are not rejected with higher p-values.

$H_{03}^{sys1}$  and  $H_{04}^{sys1}$ . Only in the second case do we fail to reject the null hypothesis (the LR statistic is 11.55, which is well under the critical value at the 1% level of significance ( $\chi^2(5) = 15.09$ )), indicating, that in Tunisia, inflation plays a significant role in the demand for money and that agricultural prices satisfy the homogeneity condition.

(Insert Table 4)

### Second subsystem

In the second subsystem, taking into account the variables included and results obtained in the first subsystem concerning the agricultural prices, the following hypothetical cointegration relations could be expected:

i) As the agricultural prices are also included in the second subsystem, and in order to check for data consistency, the first cointegration relationship would attempt to relate agricultural prices under the homogeneity restriction:

$$(\beta_1^{sys2})' Y_t : IP_t = \beta_{PP}^1 PP_t + \mu^1 + \varepsilon_{1t} \quad (11)$$

ii) The second relationship is going to be associated with an agricultural export equation for Tunisia, which would depend on the exchange rate, farm output prices and the rate of commercial openness:

$$(\beta_2^{sys2})' Y_t : AX_t = \beta_{PP}^2 PP_t + \beta_{ER}^2 ER_t + \beta_{RCO}^2 RCO_t + \mu^2 + \varepsilon_{2t} \quad (12)$$

iii) The last relationship is defined as an agricultural supply equation in which farm input and output prices, the interest rate and the rate of commercial openness are included as the main potential determinants:

$$(\beta_3^{sys2})' Y_t : AP_t = \beta_{IP}^3 IP_t + \beta_R^3 R_t + \beta_{RCO}^3 RCO_t + \mu^3 + \varepsilon_{3t} \quad (13)$$

Equations (11), (12) and (13) can be formulated in compact form as:

$$\beta'^{sys2} Y_{t-1} = \varepsilon_t \sim I(0) \text{ where } \beta'^{sys2} = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & 0 & * & 0 & * & * \\ 0 & * & 0 & 1 & 0 & * & * & * \end{bmatrix} \quad (14)$$

In order to test the restrictions on the cointegration space, a similar approach to that mentioned for the first subsystem has been followed. However, in this case, as there are three cointegration vectors, one further step has been included (Table 5). As a first step, we have carried out some tests on each individual long-run relationship, leaving

the rest unrestricted. The first hypothesis ( $H_{01}^{sys2}$ ), as mentioned above, only tries to guarantee the consistency of the data used. So, we have tested whether agricultural price homogeneity is stationary. The results from the LR test indicate that the null cannot be rejected, the same result as in the first subsystem. Three alternative hypotheses have been defined for the agricultural exports equation. The first one ( $H_{02}^{sys2}$ ) tests for a stationary relationship among agricultural exports, farm output prices, the exchange rate and the rate of commercial openness. The second one ( $H_{03}^{sys2}$ ) excludes the rate of commercial openness and includes agricultural supply. Finally, the third one ( $H_{04}^{sys2}$ ) excludes the rate of commercial openness without including any other variable. Results from the LR tests indicate that only the two first hypotheses are supported by the data. Finally, in relation to the agricultural supply equation, two alternative hypotheses have been considered. In the first one ( $H_{05}^{sys2}$ ), agricultural output is defined as a function of farm input prices, the interest rate and the rate of commercial openness. In the second one ( $H_{06}^{sys2}$ ), the rate of commercial openness is excluded from the equation. In this case, only the first hypothesis is supported by the data.

As a second step in the identification process of the long-run relationships, we have tested restrictions on two cointegration vectors taking into account the results obtained above. So, three further hypotheses have been tested. In all of them, we have maintained the price homogeneity restriction. The first one ( $H_{07}^{sys2}$ ) jointly tests  $H_{01}^{sys2}$  and  $H_{02}^{sys2}$ , whereas the second one ( $H_{08}^{sys2}$ ) tests  $H_{01}^{sys2}$  and  $H_{03}^{sys2}$  and, finally, hypothesis  $H_{09}^{sys2}$  jointly tests  $H_{01}^{sys2}$  and  $H_{05}^{sys2}$ . The results are shown in the middle part of Table 5. As can be observed, we fail to reject hypotheses  $H_{07}^{sys2}$  and  $H_{09}^{sys2}$ , indicating that we have a potential identification for the three-equation cointegration space. At the bottom of Table 5, results from jointly testing the hypotheses  $H_{01}^{sys2}$ ,  $H_{03}^{sys2}$  and  $H_{05}^{sys2}$  are shown. The null cannot be rejected, indicating that the cointegrating space is identified.

(Insert Table 5)

Finally, Table 6 shows the estimated parameters of the  $\beta$  and  $\alpha_z$  matrices corresponding to the two subsystems. In the first case, the two cointegrating vectors have been normalised by real money supply and the farm output prices. In the second subsystem, the three cointegrating vectors have been normalised by the farm output

prices, the agricultural exports and agricultural supply, respectively. All the coefficients are statistically significant and have the expected theoretical signs. The only exception is the positive sign of inflation in the money demand equation. This result, following Sriram (1999), would indicate that, in Tunisia, an expectation of increasing inflation would drive economic agents to accumulate money stock to increase nominal income. The signs associated with the rate of commercial openness in the agricultural export and production equations are also interesting. In the first case, the sign is positive as expected. In the second equation, the sign is negative, suggesting that a higher rate of commercial openness would generate increasing imports of food products in which Tunisia does not have competitive advantage (cereals, beef, vegetable oils, etc.), which negatively affects domestic production.

Juselius (1999) points out that "it is no longer possible to interpret a coefficient in a cointegrating relation as in conventional regression context....In multivariate cointegration analysis all variables are stochastic and a shock to one variable is transmitted to all variables via dynamics of the system until the system has found its new equilibrium position". So, the magnitude of the coefficients cannot be interpreted.

On the other hand, in this type of analysis, it is also convenient to consider the estimated  $\alpha_{i,j}^z$  (i indicates the row and j the column) parameters as they provide valuable information about the speed of adjustment of each variable towards the long-run equilibrium. As the relationships between macroeconomic variables and the agricultural sector are of interest for this study, let us focus on such relationships.

(Insert Table 6)

In relation to the first subsystem, and only considering the money demand and the price equations, the first conclusion is that there seems to be a feedback relationship between macroeconomic variables and the agricultural sector. In fact, any shock in the money demand generates a response of input and output prices. On the other hand, any change in the long-run relationship between agricultural prices affects both the income (Gross Domestic Product) and inflation. In relation to the price transmission mechanism, although in the long-run homogeneity holds, in the short-run the situation looks different. The  $\alpha_z$  parameters corresponding to the first cointegrating relationship indicate that input prices react quicker than output prices. This result suggests a cost-push transmission mechanism within the Tunisian agricultural sector, which is also

confirmed when observing results from the second subsystem as  $\alpha_{21} > \alpha_{11}$ ,  $\alpha_{22} > \alpha_{12}$  and  $\alpha_{23} > \alpha_{13}$ ).

In the short run, there does not seem to be a close relationship between agricultural supply and exports (the  $\alpha_{42}$  and  $\alpha_{33}$  parameters are not significant). This result would indicate that, in Tunisia, agricultural exports depend more on other factors than on the agricultural production, for example, commercial agreements (most of the exported food products are sent to the European Union and are subject to contingents) or decisions made by existing exporter lobbies in the most important export goods (olive oil, dates, citrus fruit, etc.) (see also, Allaya, 1995). In other words, agricultural policy is more oriented to supporting agricultural prices and producers' and consumers' income than to encouraging trade competitiveness. Moreover, parameter  $\alpha_{31}$  is not significant, indicating that there is no significant relationship between farm output prices and agricultural exports, which reinforces the idea of dissociation between agricultural exports and supply.

However, simply considering the magnitude of the adjustments to long-run relationships is not enough. It is also important to look at the time path of the reactions. The impulse response functions provide relevant evidence. They are analysed in the next section.

#### 4. Short-run dynamics

Once the VECM has been estimated, short-run dynamics can be examined by considering the impulse response functions (IRF). These functions show the response of each variable in the system to a shock in any of the other variables. The IRF are calculated from the Moving Average Representation of the VECM (see Lütkepohl, 1993 and Pesaran and Shin, 1998):

$$Y_t = \sum_{i=0}^{\infty} B_i \varepsilon_t$$

where matrices  $B_i$  ( $i=2, \dots, n$ ) are recursively calculated using the following expression:

$$B_n = \Phi_1 B_{n-1} + \Phi_2 B_{n-2} + \dots + \Phi_k B_{n-p}; B_0 = I_p; B_n = 0 \quad \text{for} \quad n < 0; \Phi_1 = I + \Pi + \Gamma_1;$$

and  $\Phi_i = \Gamma_i - \Gamma_{i-1}$  ( $i=2, \dots, p$ ). Following Pesaran and Shin (1998) the scaled Generalized

Impulse Response Functions (GIRF) of variable  $Y_i$  with respect to a standard error shock in the  $j^{\text{th}}$  equation can be defined as:

$$\text{GIRF}(Y_{it}, Y_{jt}, h) = \frac{e_i' B_h \Sigma e_j}{\sqrt{\sigma_{jj}}}; \quad h = 0, \dots, n$$

where  $e_s(s=i, j)$  is the  $s^{\text{th}}$  column of the identity matrix.

The GIRF are unique and do not require the prior orthogonalisation of the shocks (the reordering of the variables in the system). On the other hand, the GIRF and the orthogonalised IRF (Cholesky) coincide if the covariance matrix,  $\Sigma$ , is diagonal and  $j=1$ . Standard deviations of impulse responses are obtained following Pesaran and Shin (1998).

To analyse the short-run dynamics, as mentioned in section 2, we have first integrated the 10 variables into a full system including all the restrictions on the long-run parameters shown in the previous table. As the price transmission long-run relationship was specified in the two subsystems, only four cointegration relationships were defined.

Moreover, when defining the full system, two additional tests were carried out. The first was to check if the variables we considered as purely exogenous (R and RCO) really were. In this case, we have assumed that both variables followed a univariate autoregressive process (i.e. they are not influenced by past values of the other variables in the system) and we have tested for the significance of the appropriate  $\Gamma_i$  parameters corresponding to such equations. The value of the statistic was 31.02, which was under the critical value ( $\chi^2_{17} = 33.41$ ) at the 1% level of significance). Second, all the adjustment coefficients ( $\alpha_z$  parameters), which were non-significant in Tables 3 and 4, were restricted to zero. The test indicated that it was not possible to reject the null (the LR statistic was 58.92, which was under the critical value of a  $\chi^2(43)=66.95$ , at the 1% significance level). The final estimated model is given by:

$$\beta'Y_i = \begin{bmatrix} 1,000 & -5,339 & -1,413 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,014 & 0,000 & 0,586 \\ 0,000 & 0,000 & 0,000 & 1,000 & -1,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,207 \\ 0,000 & 0,000 & 0,000 & 0,748 & 0,000 & 1,000 & 0,000 & -1,067 & 0,000 & -0,303 & -1,176 \\ 0,000 & 0,000 & 0,000 & 0,000 & 0,871 & 0,000 & 1,000 & 0,000 & 0,074 & 0,894 & -0,277 \end{bmatrix} \begin{Bmatrix} M \\ P \\ GDP \\ PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \\ Const. \end{Bmatrix}$$

$$\alpha = \begin{bmatrix} 0,0000 & 0,0000 & 0,0000 & 0,0000 \\ 0,0333 & 0,0141 & 0,0000 & 0,0000 \\ 0,0516 & 0,0252 & 0,0000 & 0,0000 \\ 0,0521 & 0,0324 & 0,0128 & 0,0117 \\ 0,1786 & 0,1035 & 0,0464 & 0,0409 \\ 0,0000 & 0,0000 & 0,1454 & 0,0000 \\ 0,0000 & 0,1555 & 0,0000 & 0,0495 \\ 0,0000 & 0,0000 & 0,0000 & 0,0000 \\ 0,0000 & 0,0000 & 0,0000 & 0,0000 \\ 0,0000 & 0,0000 & 0,0000 & 0,0000 \end{bmatrix} \begin{Bmatrix} M \\ P \\ GDP \\ PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \end{Bmatrix}$$

As can be observed, the long-run structural parameters are exactly the same than those shown in Table 6, which guarantees the consistency of the methodological approach followed here<sup>14</sup>. As 100 impulse-response functions are obtained, we will analyse only the estimated impulse responses of agricultural variables to a shock in the main variables of the system. Significant responses are marked with a circle. In general terms, most of the responses are not significant although they show the expected signs. In any case, as we have annual data, we cannot expect responses longer than one or two years. Figure 1 shows the responses to a shock in the real quantity of money, (through an unexpected increase in the nominal quantity of money). As expected, an expansive monetary policy positively affects inflation, although the effect is only significant during the second year after the shock. Moreover, increasing access to credits stimulates economic growth (the GDP increase) as well as agricultural exports. The effect on farm output prices is

<sup>14</sup> As we were interested in analysing the full set of interrelationships between macroeconomic and agriculture variables, the estimation of a single system with the 10 variables would have provided more efficient estimates. However, as only 35 observations were available for each series, the estimation of a full system, as mentioned in Section 2, was not possible due to the lack of degrees of freedom. The methodological approach followed here allowed us to analyse all possible interrelationships without a significant loss of efficiency. Moreover, when comparison was possible, the impulse response functions derived from partial models were not significantly different than those shown in this paper.



positive but non significant. The aim of the Tunisian agricultural policy is to support farmers' income through reasonably high intervention prices although keeping inflation controlled. Agricultural prices are not allowed to increase over expected inflation in order to guarantee consumer access to basic foods. Limited increases of agricultural prices do not stimulate either agricultural production or the demand for inputs. As a result, the impact on input prices is not significant. Finally, following the Keynesian theory, an expansive monetary policy induces exchange rate depreciation, which, leads to an increase of agricultural exports.

(Insert Figure 1)

Figure 2 shows the responses to a shock in farm output prices of the most relevant variables within the system. In general terms, agricultural variables do not have any effect on macroeconomic variables and, so they are not included in the Figure. Two main results are found. First, a positive increase of producer prices generates a positive response of agricultural production. The response is significant for two years. Second, it also generates an immediate positive response of input prices due to the increasing derived input demand as a consequence of production increases. The magnitude of this response is higher than in the case of output prices, which is consistent with the comments in Section 3 about the  $\alpha_z$  parameters and also with the trend followed by both price series during the analysed period. El Abassi (1995) and Allaya (1995) explain this trend by the intensification process that the Tunisian agriculture has suffered during the last two decades and the progressive elimination of subsidies addressed to fertilizers and other intermediate inputs. On the other hand, the effect on inflation is positive but not significant. As mentioned above, the government has traditionally controlled agricultural price increases to be compatible with the inflation rate in Tunisia. Finally, the effect on agricultural exports is not significant either. Prices are not the main source of competitiveness for Tunisian agricultural exports as they are mainly subject to contingents. Moreover, the traditional policy used by the Tunisian Government to promote exports has been via the exchange rate.

(Insert Figure 2)

Responses to a shock in the farm input price are shown in Figure 3. Some interesting results are found. The first is that responses of farm output prices are of lower magnitude than those of input prices, a result which is consistent with the

empirical evidence provided by El Abassi (1995). Moreover, the response of output prices is only significant two years after the initial shock. It seems that the public authorities increase intervention prices as a consequence of increasing production costs and that it takes another season for producers to adapt to the new situation. In any case, and taking into account all the results found in this paper on price transmission, we can conclude that a cost-push transmission mechanism prevails in the Tunisian agro-food sector. This result can be also confirmed by the significant response of inflation to increasing input prices. Finally, a positive shock in production costs reduces agricultural output the following year, which would also contribute to explaining the significant increase of output prices during the second year after the shock. However, the effect on agricultural exports is not significant.

(Insert Figure 3)

A positive shock in the exchange rate generates an immediate positive reaction in agricultural exports (Figure 4), confirming the idea we have mentioned above about the exchange rate as an important short-run determinant of the competitiveness of agricultural exports. However, in the long-run the effect is not significant. This results is outlined also in Ben Said (1994) and Allaya (1995) who concluded that after the devaluation of the Tunisian dinar that took place with the implementation of the Structural Adjustment Program the growth rate of Tunisian agricultural exports was lower than that of the exchange rate. Finally, as can be also observed in Figure 4, the effect on the rest of variables is negligible.

(Insert Figure 4)

## 5. Concluding remarks

The aim of this paper is to apply recent developments in the econometric analysis of time series to the study of relationships between macroeconomic variables and the agricultural sector in Tunisia. Results from this study suggest a number of points. The first is that it is interesting to distinguish between long-run and short-run analyses. The long-run analysis is usually associated with structural relationships and it is in this context that theoretical restrictions have to be tested. The short-run analysis is also important for policy analysis as it gives an idea of the magnitude and time path of the reactions of economic variables to deviations from long-run relationships. However, the

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short-run responses of variables to shocks have to be calculated with the aid of theoretically based long-run economic restrictions.

The study has shown that changes in agricultural variables have no significant effects on macroeconomic variables. Only shocks in agricultural prices have an effect on inflation. The main source of responses of the agricultural sector (mainly agricultural output and exports) is changes in the monetary policy and, more precisely, in money supply, which is consistent with how monetary policy is instrumented in Tunisia.

Agricultural prices responses to macroeconomic shocks are not very significant as an indication of the degree of government intervention in Tunisia. In the case of agricultural exports, responses are larger if they are generated by changes in the exchange rate than if they are generated by changes in output prices, which is an indication that macroeconomic variables have to be taken into account when designing agricultural policy. In the same context, agricultural supply is quite inelastic but reacts more to changes in capital costs than to changes in input or output prices. To conclude, it has to be said that the results presented in this paper depend on the variables and sample period chosen. Further analysis, including other variables and an extended sample period, could be conducted in the future.

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For Peer Review



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APPENDIX

Variable	Symbol	Source	Units
Exchange rate	ER	International Monetary Fund (IMF)	Tunisian Dinars per US dollar
Interest rate	R	Statistiques financières. Banque Centrale de Tunisie.	Percentage
Money Supply	M	International Monetary Fund (IMF)	Million dinars
Consumer Price Index	P	International Monetary Fund (IMF)	Index (Basis 100 = 1990)
Gross Domestic Product	GDP	International Monetary Fund (IMF)	1990 Million dinars
Rate of Commercial Openness	RCO	International Monetary Fund (IMF)	Percentage
Farm output prices	PP	Institut National de la Statistique. Ministère du développement économique. Tunisie.	Index (Basis 100 = 1990)
Farm input prices	IP	Institut National de la Statistique. Ministère du développement économique. Tunisie.	Index (Basis 100 = 1990)
Agricultural exports	AX	FAO.	1990 Million dinars
Agricultural Output	AP	Institut National de la Statistique. Ministère du développement économique. Tunisie.	1990 Million dinars



Table 1. Results from Lanne et al. (2003) unit root tests with structural breaks.

Variable	Break point ( $T_B$ )	With trend	Break point ( $T_B$ )	Whithout trend
ER	1981	-1.82	1980	-0.33
R	1981	-1.58	1981	-1.50
M	1974	-1.81	1974	-1.56
P	1996	-2.12	1980	-1.85
GDP	1982	-2.68	1982	-2.12
RCO	1974	-1.44	1976	-2.42
PP	1982	-1.28	1982	-2.17
IP	1986	-2.59	1986	-2.56
AX	1972	-1.77	1976	-2.02
AP	1971	-2.21	1995	-0.91
CV (5%) <sup>a</sup>		-3.15		-2.99

<sup>a</sup> Critical values are from Lanne et al. (2002) (Table 2, T=50).

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Table 2. Results from cointegration tests

First subsystem $Y' = \{M, P, GDP, PP, IP, R\}$ <sup>1</sup>							
$H_0: r$	$H_a: p-r$	$\lambda$ -max.	Trace	Critical values $\lambda$ -max <sup>2</sup>		Critical values Trace <sup>2</sup>	
				(10%)	(5%)	(10%)	(5%)
0	5	46,88	124,76	34,99	37,48	82,17	86,58
1	4	37,72	77,88	29,01	31,48	59,07	62,75
2	3	20,92	38,15	22,98	25,54	39,12	42,40
3	2	11,37	19,23	16,74	18,88	22,76	25,23
4	1	7,85	7,85	10,50	12,45	10,50	12,45
Second subsystem $Y' = \{PP, IP, AX, AP, ER, R, RCO\}$ <sup>1</sup>							
$H_0: r$	$H_a: p-r$	$\lambda$ -max.	Trace	Critical values $\lambda$ -max <sup>2</sup>		Critical values Trace <sup>2</sup>	
				(10%)	(5%)	(10%)	(5%)
0	5	42,84	123,36	37,81	40,57	92,93	97,57
1	4	34,91	80,15	32,00	34,69	67,83	72,15
2	3	28,42	47,06	26,08	28,49	45,89	49,43
3	2	16,17	23,85	19,67	21,92	27,58	30,46
4	1	7,68	7,68	13,21	15,27	13,21	15,27

<sup>1</sup> See the Appendix for definitions of the variables

<sup>2</sup> Critical values are taken from Pesaran et al. (2000).

Table 3. Estimated  $\beta$  and  $\alpha_z$  parameters for both subsystems

$\beta$ parameters <sup>1</sup>						
First subsystem (r=2)			Second subsystem (r=3)			
Variable	Vector 1	Vector 2	Variable	Vector 1	Vector 2	Vector 3
M*	1,000	-0,130	PP*	-2,471	-1,112	1,000
P*	-5,157	-0,764	IP*	-1,468	-3,360	-0,079
GDP*	-1,366	0,101	AX*	1,000	-0,533	0,099
PP*	-0,381	1,000	AP*	-1,068	1,000	-0,312
IP*	-0,283	-0,052	ER*	1,011	4,125	0,029
R*	0,017	-0,006	R*	0,079	0,026	0,015
Constant*	8,916	-0,231	RCO*	-1,210	-3,552	-0,105
			Constant*	2,188	-8,913	1,961
$\alpha_z$ parameters						
DM	-0,166	0,762	DPP	0,119	-0,015	-0,137
DP	0,175	-0,131	DIP	0,374	0,045	-0,666
DGDP	0,213	0,541	DAX	-0,431	0,100	-2,183
DPP	-0,068	-0,015	DAP	0,218	0,016	1,253
DIP	-0,292	-0,159	DER	-0,152	-0,018	-0,455

<sup>1</sup> An \* indicates that the parameter is significant at the 5% level of significance (critical values are 5.99 and 7.81 for the first and second subsystems, respectively)

Table 4. Hypothesis restriction tests on the cointegration vectors in the first subsystem <sup>a</sup>

Hypotheses on a single cointegration vector		
Hypothesis formulation	Statistic	Critical Value (1%)
$H_{0i} : \beta_i^{sys1} = (\beta_1, \Phi) = (H_1\phi, \Phi)$		
$H_{01}^{sis1} : \beta_1^{sys1} Y_t = \begin{bmatrix} 1 & * & -1 & 0 & 0 & * & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(2) = 8,77$	9,21
$H_{02}^{sis1} : \beta_2^{sys1} Y_t = \begin{bmatrix} 1 & 0 & -1 & 0 & 0 & * & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(3) = 12,37$	11,34
$H_{03}^{sis1} : \beta_3^{sys1} Y_t = \begin{bmatrix} 1 & * & * & 0 & 0 & * & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 5,58$	6,63
$H_{04}^{sis1} : \beta_4^{sys1} Y_t = \begin{bmatrix} 0 & 0 & 0 & 1 & -1 & 0 & * \\ * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(4) = 10,12$	13,28
Hypotheses on the full system		
Hypothesis formulation	Statistic	Critical value (1%)
$H_{0i} : \beta_i^{sys1} = (\beta_1, \beta_2) = (H_1\phi_1, H_2\phi_2)$		
$H_{05}^{sis1} : \beta_5^{sys1} Y_t = \begin{bmatrix} 1 & * & -1 & 0 & 0 & * & * \\ 0 & 0 & 0 & 1 & -1 & 0 & * \end{bmatrix} Y_t$	$\chi^2(6) = 35,16$	16,81
$H_1' = \begin{bmatrix} 1 & 0 & -1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$ and $H_2' = \begin{bmatrix} 0 & 0 & 0 & 1 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$		
$H_{06}^{sis1} : \beta_6^{sys1} Y_t = \begin{bmatrix} 1 & * & * & 0 & 0 & * & * \\ 0 & 0 & 0 & 1 & -1 & 0 & * \end{bmatrix} Y_t$	$\chi^2(5) = 11,55$	15,09
$H_1' = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$ and $H_2' = \begin{bmatrix} 0 & 0 & 0 & 1 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$		

<sup>a</sup> Y' = {M, P, GDP, PP, IP, R}. An \* indicates that the coefficient is unrestricted.

Table 5. Hypothesis restriction tests on the cointegration vectors in the second subsystem<sup>a</sup>

Hypotheses on a single cointegration vector		
Hypothesis formulation	Statistic	Critical value (1%)
$H_{0i} : \beta_i^{\text{sys2}} = (\beta_1, \Phi) = (H\phi, \Phi)$		
$H_{01}^{\text{sis2}} : \beta_1^{\text{sys2}} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(4) = 7,87$	13,28
$H_{02}^{\text{sis2}} : \beta_2^{\text{sys2}} Y_t = \begin{bmatrix} * & 0 & 1 & 0 & * & 0 & * & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 0,01$	6,63
$H_{03}^{\text{sis2}} : \beta_3^{\text{sys2}} Y_t = \begin{bmatrix} * & 0 & 1 & * & * & 0 & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 0,50$	6,63
$H_{04}^{\text{sis2}} : \beta_4^{\text{sys2}} Y_t = \begin{bmatrix} * & 0 & 1 & 0 & * & 0 & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(2) = 9,83$	9,21
$H_{05}^{\text{sis2}} : \beta_5^{\text{sys2}} Y_t = \begin{bmatrix} 0 & * & 0 & 1 & 0 & * & * & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(1) = 6,06$	6,63
$H_{06}^{\text{sis2}} : \beta_6^{\text{sys2}} Y_t = \begin{bmatrix} 0 & * & 0 & 1 & 0 & * & 0 & * \\ * & * & * & * & * & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(2) = 10,70$	9,21
Hypotheses on two cointegration vectors		
Hypothesis formulation:	Statistic	Critical value (1%)
$H_{0i} : \beta_i^{\text{sys2}} = (\beta_1, \beta_2, \Phi) = (H_1\phi_1, H_2\phi_2, \Phi)$		
$H_{07}^{\text{sis2}} : \beta_7^{\text{sys2}} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & 0 & * & 0 & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(5) = 8,38$	15,09
$H_{08}^{\text{sis2}} : \beta_8^{\text{sys2}} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & * & * & 0 & 0 & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(5) = 17,23$	15,09
$H_{09}^{\text{sis2}} : \beta_9^{\text{sys2}} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ 0 & * & 0 & 1 & 0 & * & * & * \\ * & * & * & * & * & * & * & * \end{bmatrix} Y_t$	$\chi^2(5) = 10,07$	15,09
Hypotheses on the three cointegration vectors		
Hypothesis formulation	Statistic	Critical value (1%)
$H_{0i} : \beta_i^{\text{sys2}} = (\beta_1, \beta_2, \beta_3) = (H_1\phi_1, H_2\phi_2, H_3\phi_3)$		
$H_{010}^{\text{sis2}} : \beta_{10}^{\text{sys2}} Y_t = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & * \\ * & 0 & 1 & 0 & * & 0 & * & * \\ 0 & * & 0 & 1 & 0 & * & * & * \end{bmatrix} Y_t$	$\chi^2(6) = 15,23$	16,81
$H'_1 = \begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}; H'_2 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}; H'_3 = \begin{bmatrix} 0 & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}$		

<sup>a</sup>  $Y' = \{\text{PP, IP, AX, AP, ER, R, RCO}\}$ . An \* indicates that the coefficient is unrestricted.

Table 6. Estimated  $\beta$  and  $\alpha_z$  matrices under long-run identification for both subsystems

First subsystem

$$\beta' = \begin{bmatrix} 1,000 & -5,339 & -1,413 & 0,000 & 0,000 & 0,014 & 0,586 \\ 0,000 & 0,000 & 0,000 & 1,000 & -1,000 & 0,000 & 0,207 \end{bmatrix} \times \begin{bmatrix} M \\ P \\ GDP \\ PP \\ IP \\ R \\ Const. \end{bmatrix}$$

Standard deviations for  $\beta$  parameters: (0,413), (0,026), (0,007), (0,231), (0,471)

$$\alpha_z = \begin{bmatrix} -0,123 & -0,014 \\ 0,156 & 0,034 \\ 0,204 & 0,066 \\ 0,096 & -0,023 \\ 0,314 & 0,076 \end{bmatrix}$$

t-ratios for  $\alpha_z$  parameters: (-1,827), (-0,522), (4,343), (3,752), (3,034), (3,883), (1,945), (-2,787), (1,978), (2,691)

Likelihood Ratio Statistic= 11,55  
Critical value (1%) = 15,09

Second subsystem

$$\beta' = \begin{bmatrix} 1,000 & -1,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,000 & 0,269 \\ 0,748 & 0,000 & 1,000 & 0,000 & -1,067 & 0,000 & -0,303 & -1,176 \\ 0,000 & 0,871 & 0,000 & 1,000 & 0,000 & 0,074 & 0,894 & -0,277 \end{bmatrix} \times \begin{bmatrix} PP \\ IP \\ AX \\ AP \\ ER \\ R \\ RCO \\ Const. \end{bmatrix}$$

Standard deviations for  $\beta$  parameters: (0,058), (0,384), (0,083), (0,194), (0,823), (0,080), (0,178), (0,862)

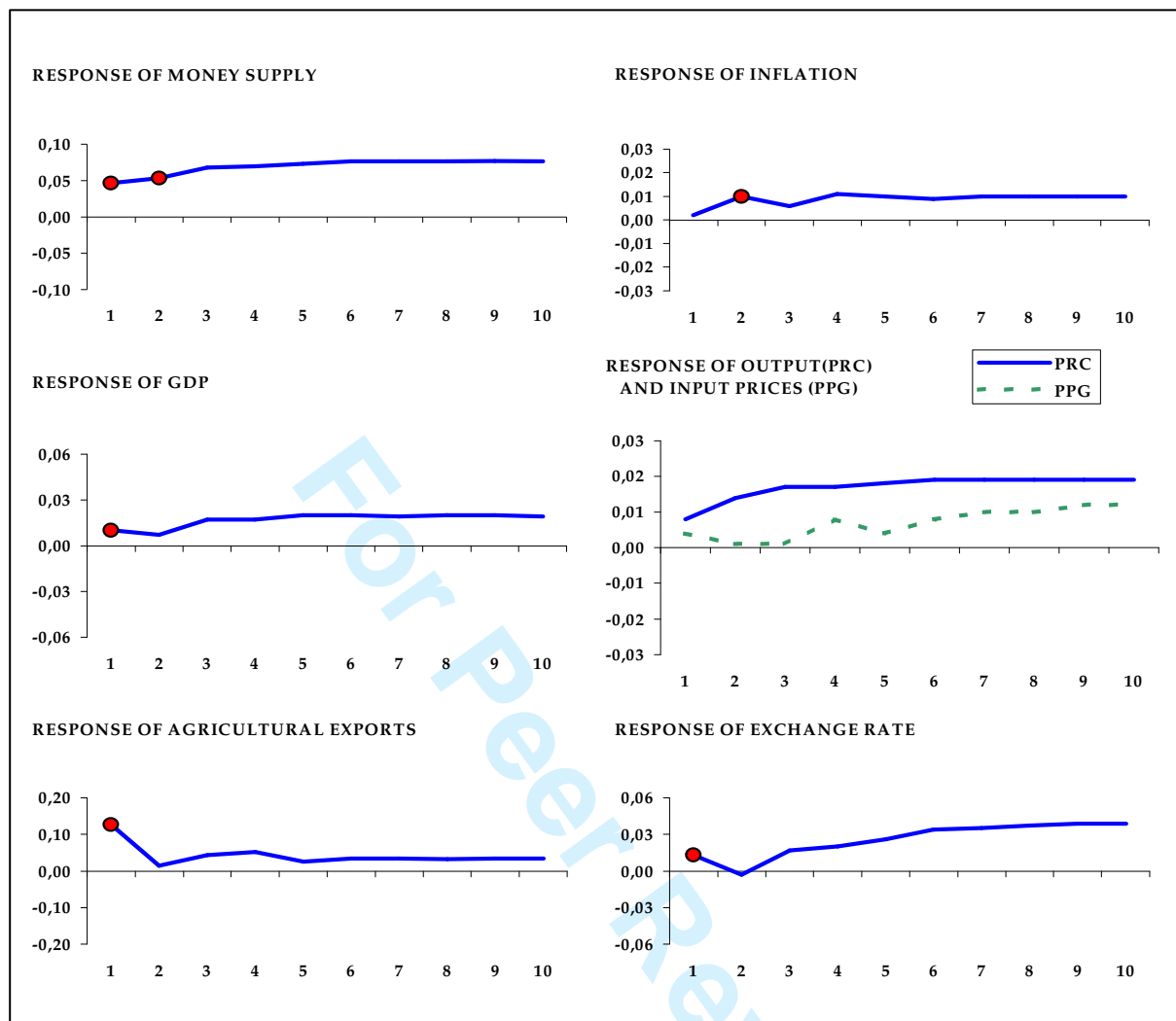
$$\alpha_z = \begin{bmatrix} -0,146 & 0,079 & -0,049 \\ 0,260 & 0,109 & -0,361 \\ -0,217 & -0,363 & 0,400 \\ 0,145 & 0,260 & -0,128 \\ -0,048 & -0,028 & 0,003 \end{bmatrix}$$

t-ratios for  $\alpha_z$  parameters: (-2,279), (2,781), (-2,083), (4,014), (4,328), (-3,871), (-0,832), (-3,898), (0,786), (2,356), (1,323), (-1,932), (-0,723), (-1,198), (0,873)

Likelihood Ratio Statistic = 15,23  
Critical Value (1%) = 16,81

Note: Values in parentheses correspond to standard deviations, in the case of the  $\beta$  parameters, and to t-ratios, in the case of the  $\alpha_z$  parameters

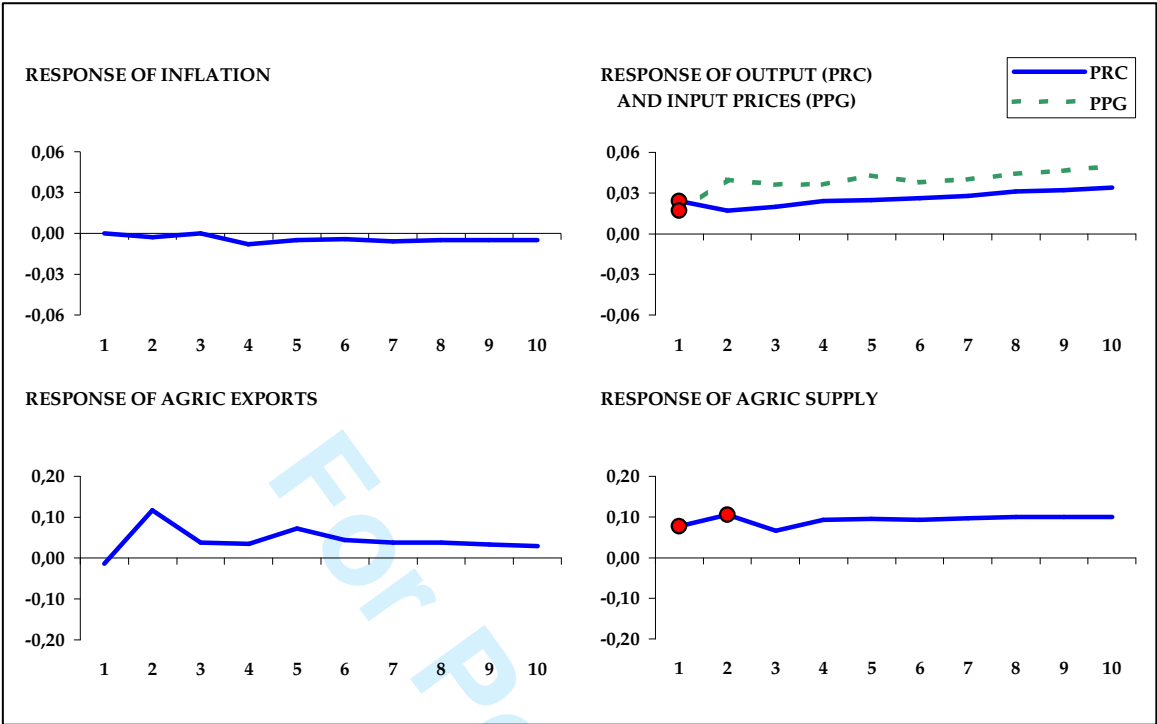
Figure 1. Responses to a shock in the money supply



Note: Significant responses at the 5% level of significance are marked with a circle



Figure 2. Responses to a shock in farm output prices



Note: Significant responses at the 5% level of significance are marked with a circle

Figure 3. Responses to a shock in farm input prices

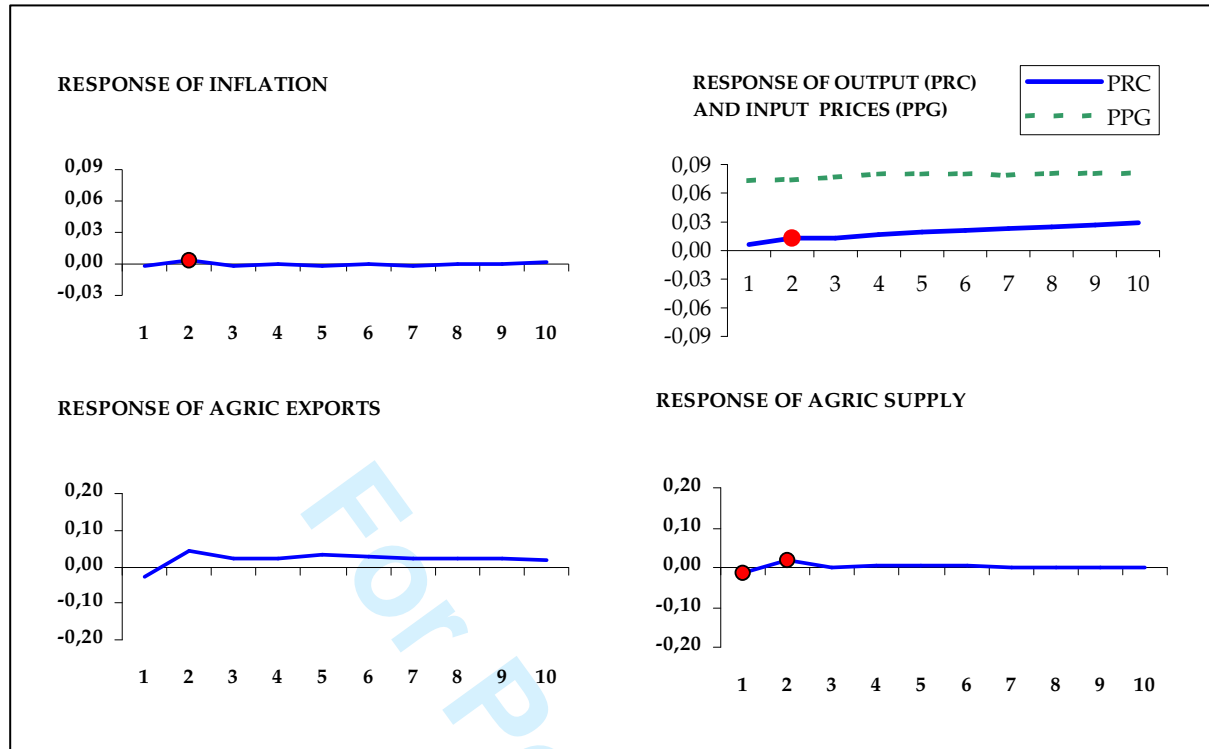
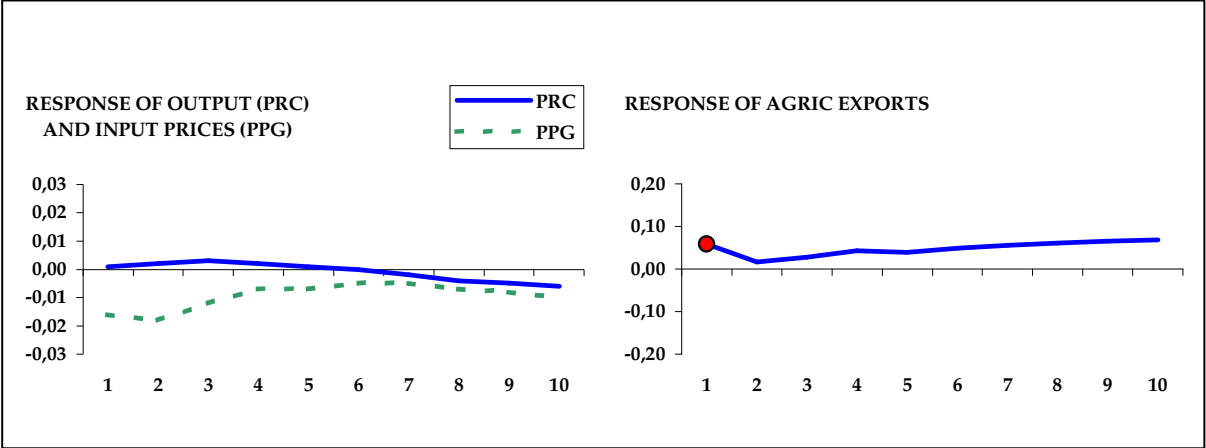


Figure 4. Responses to a shock in the exchange rate



Note: Significant responses at the 5% level of significance are marked with a circle